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ABSTRACT

This paper builds on the work of Deaton and Laroque (2003) by formulating a nonlinear model of commodity prices. The paper makes three distinct contributions. First, a nonlinear model is constructed that explains long-run dynamics of commodity price behavior; secondly, more recent data is employed by updating the price, income and production indices; and finally advanced econometric techniques are adopted in order to investigate whether there is empirical evidence to support the theoretical underpinnings of the nonlinear model. Higher power tests broadly reverse the empirical findings of Deaton and Laroque's (2003) model of commodity prices, lending support to the underlying theory proposed in this paper. Tests for cointegration provide evidence that the long-run relationship between world commodity production and world income for key commodities such as sugar, copper, and tin may be better explained by non-linear behavior.

Keywords: Commodity prices; Lewis Model; Cointegration, ESTAR.

JEL classification codes: E3; F1; O1

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1. Introduction

The dynamics of commodity price behavior have received considerable interest from empirical researchers. Developing countries whose export earnings are heavily reliant on one or few primary commodities are extremely sensitive to the dynamics of commodity prices. Cashin *et al.* (2002) emphasize the reliance of less-developed countries (LDCs) on the international trade of one or a few key primary commodity exports, such as Ethiopia (coffee), Zambia (copper), and Mauritius (sugar). Developing countries can experience major problems when the prices of their exports (primary commodities) are found to decline over a period of time in relation to the price of their imports (manufactured goods). The dynamic behavior of commodity prices is influenced by the underlying stochastic or deterministic nature of the trends of commodity prices. An understanding of the stochastic process of primary commodity prices is therefore essential for policy makers in developing countries.

There is a vast amount of research that has addressed whether commodity prices contain stochastic trends or are stationary processes. The studies have been empirical, which are motivated by unit root tests. This approach has delivered a large volume of research with mixed results as to whether unit roots do exist in commodity prices. Previous literature has focused on the long-run decline in the terms of trade, or the Prebisch–Singer Hypothesis, between the primary goods exports of developing countries and the manufactured goods exports of industrialized countries [see, for example, Ghoshray (2011), Harvey *et. al.* (2010), Balagtas and Holt (2009), Kellard and Wohar (2006)].

A theory which Deaton (1999) is more convinced by is a model which argues that developing country workers who work at a fixed subsistence wage level do not actually profit from any technological changes made in the sector in which they work, whereas for “a wheat farmer in Canada, whose wage is set in the industrial labor markets of North America where the aggregate supply of labor is limited, …real wages can therefore rise in response to technological change” (Deaton, 1999, pp.29-30).

Deaton and Laroque (2003) (DL henceforth) depart from the use of speculative storage as a short-run determinant of commodity prices and instead focus on supply and demand factors in the long run. They develop a time series model of commodity prices, which provides the long-run dynamics of commodity price behavior, and in turn leads to the world supply of commodities changing in line with world demand (proxied by world GDP). While DL acknowledge that “because demand is more highly autocorrelated, it is a good candidate to explain autocorrelation” (DL, pp.290), more attention is paid to supply-side characteristics. DL uses the Lewis model (1954) to outline the supply-side determinants of commodity prices. Core to the Lewis model of economic development is the infinite labour supply working at a subsistence wage level. In a commodities sense, this represents a key feature of DL because the model assumes that "commodity supply is infinitely elastic in the long run" (*ibid.*, pp.290). A follow-on from Deaton and Laroque's 1996 paper is the use of p (the current price) and p^* (the critical price). In DL, commodity supply growth is a function of how far p (the commodity price) deviates from p^* (the long-run commodity supply price). DL assumes commodity price stationarity around the supply price and cointegration between world supply and world income. The assumption of long supply lags, which is common for certain agricultural commodities, allows DL to argue that "the price process....... in the short run is driven by fluctuations in world income..." (DL, pp.290). However, once supply lags are accounted for, prices are stationary in the long run.

The empirical results in DL, however, do not conform to the theoretical underpinning of the Lewis model. DL make use of standard linear unit root and cointegration tests which do not provide support to their theoretical model. In this paper we suggest that certain assumptions could be relaxed or changed to provide a similar but alternative model that can be better represented by the empirics. Consequently, this paper aims to extend the work of DL in addressing how well the Lewis (1954) model can be adapted to explain commodity prices.

Bond (1983) highlights certain characteristics of commodities in Sub-Saharan Africa which suggest that output may not be responsive to changes in price (*ibid.*, pp.705). It is the conventional wisdom that subsistence farmers display risk averse behavior and attach greater value to leisure than using their labour to increase output. Another explanation is that African farmers have “income targets”, so a small rise in producer prices suggests that farmers need not produce as much in order to reach the same income target (*ibid.*, pp.705).

This paper looks to contribute to the existing literature by extending several areas of DL’s work, namely in terms of the extension of the economic model, the timeliness of datasets used for empirical estimations and the sophistication of the econometric techniques used. The extension of DL’s version of the Lewis model of commodity prices is carried out with regards to the commodity supply function by introducing the non-linearity in the responsiveness of commodity supply to fluctuations in commodity prices. Therefore, we assume that supply is more responsive to a large deviation in price from its long-run equilibrium compared to the case where there is a minimal price deviation. In terms of the data, this paper uses twenty more years’ worth of commodity price data (that is, cocoa, coffee, rice, sugar, copper and tin) taken from the Grilli-Yang index. Second, the paper addresses DL’s difficulty that “long-run production and income data are a great deal harder to come by” (DL, p.297) beyond 1987. Indeed, the current study builds on the world commodity production data accumulated by DL and introduces other sources for the post-1987 period data. This paper also uses an alternative series for world income in lieu of that provided by DL (p. 307-9) so that the data spans a longer time period to coincide with the longer time series for prices and production. Finally, we apply more recent econometric estimation techniques when testing for stationarity of commodity prices, production and world income and also when testing for evidence of cointegration between commodity production and world income.

The paper comprises six sections, which are detailed as follows: Section 2 details the underlying theory of the Lewis Model derived by DL and our modifications; 3 provides a description of the updated data used in this study; Section 4 presents the econometric methodology; Section 5 describes the empirical findings and the policy implications; and finally Section 6 concludes.

2. Economic Model

The Lewis model has been founded on a stylized dualistic developing economy which comprises an agricultural sector and an industrial sector. The model rests on the underlying assumption of there being an infinite supply of labourers within the agricultural sector who are paid a fixed, low and exogenously-determined subsistence wage. The Lewis model emphasizes that at the subsistence wage, there is an excess supply of labour and that the excess supply is sufficiently large, so that employers need not concern themselves with the threat of wages being bid up by labour force shortages. With regards to the primary commodities sector, this implies that commodity-export dependent countries will struggle to experience a rise in their commodity prices – furthermore, the advent of greater and faster technical progress may even cause primary commodity prices to fall further. The industrial sector, however, must be prepared to pay a mark-up on the agricultural sector subsistence wage in order to attract labourers.

DL make use of a time-series approach to the Lewis model. In their model, the behavior of commodity prices stems from the interaction of “an integrated (trending) demand process against a supply function that is infinitely elastic in the long run but not the short run” (DL, p.291). While the Lewis model finds its origins in the agricultural sector, DL also link the model to the situation faced by workers in the metals and minerals industry. The justification is that “if the extraction costs are the main component of the observed price, and if these costs are the wages of subsistence workers employed to do the extraction” (DL, p.292), then commodity price rents will hinge on the proportion of these costs relative to total costs of production.

DL develop a partial-equilibrium analysis of the price determination of an internationally traded commodity. Demand is assumed to have constant elasticity and be a function of world income (world GDP) and of the world price (DL, p.293). The natural logarithm of world income levels is characterized by a non-stationary $I(1)$ process, that is, world income levels are found to be non-stationary in levels and stationary in first-differences. Equation (1) below shows that the natural logarithm of quantity of commodities demanded (d_t) is a function of logged commodity prices (p_t) and logged levels of world income (y_t) .

$$d_t = Ay_t - Bp_t + k \tag{1}$$

In equation (1), DL assume $k > 0$, $B > 0$ and $A > 0$. In contrast to DL, we have chosen to drop the error term ξ_t^d from the demand function, where DL assume that ξ_t^d is unobservable, $I(0)$ random variable.[1]

DL specify the supply function by equation (2) given below. The change in commodities supplied (s_t) is a linear function of the excess of price of the commodity (p_t) over its marginal cost p^*, and "an unobservable, stationary, $I(0)$ random variable" ξ_t^s (*ibid.*).

$$s_t = s_{t-1} + D(p_t - p^*) + \xi_t^s \tag{2}$$

In an effort to span the agricultural as well as the metals and minerals industry, p^* can represent the "marginal cost of production on marginal land" (*ibid.*) with regards to the former, or the "marginal cost of extraction for a mineral" (*ibid.*) with regards to the latter. Supply increases when price is above marginal cost and is described in (2) with $D > 0$. We assume that the supply shock (ξ_t^s) is a function of permanent (more

[1] We argue that the demand function is unlikely to be stochastic. Change in demand of world consumption of commodities appear to be more long-term as there is more focus on changes in consumer preferences. Consumer tastes and preferences imply that changes in world demand represent a lengthier process. Hence, world demand fluctuations are more rarer than supply fluctuations.

likely in minerals markets) and transitory shocks (weather-related in agricultural markets). Labour costs and productivity can fall into either category.

However, in this paper we intend to modify the supply function given by equation (2). Following the argument put forward by Bond (1983) we feel that the change in supply being responsive at a constant proportion of the deviation of price from marginal cost is an unrealistic assumption. One would expect the supply changes to be very small when price is marginally higher than marginal cost. However the change in supply is more responsive when price deviates to a greater extent from marginal cost. Moreover, suppliers, especially in developing countries, do not have precise information regarding the market. It is likely that they will base their expectations of price in the current period on what prices were in the previous period. The supply function given by (2) can be modified to allow the supply function to take the following form:

$$s_t = s_{t-1} + D(p_{t-1} - p*)\left[1 - e^{-\theta(P_{t-1} - P*)^2}\right] + \xi_t^s \tag{3}$$

Following the assumption in DL that markets clear, we can use (1) and (3) [see Appendix A.I for details] to arrive at the reduced form price equation as follows:

$$\Delta\tilde{P}_t = -\frac{D}{B}\tilde{P}_{t-1}\left[1 - e^{-\theta\tilde{P}_{t-1}^2}\right] + \frac{A}{B}\Delta y_t - \frac{1}{B}\left[\xi_t^s\right] \tag{4}$$

where $\widetilde{P}_t = P_t - P^*$. Equation (4) shows that in the short run, price will respond to fluctuations in demand and supply. If prices deviate further from their steady-state inter-temporal equilibrium, given by $A\phi/B$ where $\phi = E(\Delta y_t)$ is the mean growth rate of income, then prices will show signs of adjustment. Equation (4) shows that the commodity price should be stationary with nonlinear adjustment, fluctuating around its long-run value of $A\phi/B$, where $\phi = E(\Delta y_t)$ is the mean growth rate of income. Similar to DL, the price series has no long-run trend. In other words, if we were to conduct unit root tests on the price series, we would expect to obtain stationary processes with nonlinear adjustment.

DL, in their paper derive the stochastic supply process, which can be expressed as a cointegrating relation between income and supply (the latter being proxied by production of that commodity). We seek to estimate the same model as DL, that is, testing for the long-run relationship or cointegration of income and production; however, we also allow for nonlinear adjustment to any deviation from this long-run relationship. This is motivated by the supply function given by (3) that describes a nonlinear adjustment process. Combining (1) and (3), and after some algebraic manipulation [see Appendix A.II], we can obtain the following stochastic function:

$$\Delta s_t = D\left[-\frac{1}{B}d_{t-1} + \frac{A}{B}y_{t-1} + \frac{k}{B} - P^*\right]\left[1 - e^{-\theta\left(-\frac{1}{B}d_{t-1} + \frac{A}{B}y_{t-1} + \frac{k}{B} - P^*\right)^2}\right] + \xi_t^s \qquad (5)$$

Following from DL that markets clear in every period we obtain;

$$\Delta s_t = D\left[-\frac{1}{B}s_{t-1} + \frac{A}{B}y_{t-1} + h\right]\left[1 - e^{-\theta\left(-\frac{1}{B}s_{t-1}+\frac{A}{B}y_{t-1}+h\right)^2}\right] + \xi_t^s \qquad (6)$$

where $h = \left(k/B - P^*\right)$. For simplicity, we assume the demand shock variable ξ_t^d to be zero on the grounds that it is a stationary, unobservable, I(0) random variable. The stochastic supply function can be now represented as follows:

$$\Delta s_t = -\frac{D}{B}\left[\left(s_{t-1} - Ay_{t-1} - Bh\right)\right]\left[1 - e^{-\theta\left\{-\frac{1}{B}\left(s_{t-1}-Ay_{t-1}-Bh\right)\right\}^2}\right] + \xi_t^s \qquad (7)$$

This long-run value of the price guarantees that, over the long run, supply increases at the same rate as world income so that, as shown in (7), supply is cointegrated with world income. This is verified by the nonlinear error correction model that we formulate in equation (7) which shows that any deviation from the long run equilibrium between supply (that is, production) and income is corrected in a nonlinear fashion.

3. Data description

There are three types of data which are used in the empirical section of this paper: world gross domestic product (GDP) levels, real commodity prices for six different commodities (cocoa, coffee, rice, sugar, copper, and tin), and their world production levels. This paper attempts to observe and carry out empirical work on an extended sample period (1900-2008) so as to incorporate the price, supply, and world income fluctuations experienced over the last twenty years.

3.1. Updating real non-fuel primary commodity prices

The six non-fuel primary commodity price series used in this paper have been taken from the well-known and much-used Grilli and Yang (1988) index of commodity prices. Grilli and Yang (GY henceforth) constructed "a U.S. dollar index of prices of twenty-four internationally-traded non-fuel commodities, beginning in 1900" (GY, 1988, p.3) in order to address the data inadequacies of commodity price indices of the time. The Grilli and Yang (1988) commodity price index (GYCPI) is "base-weighted, with 1977-79 values of world exports of each commodity used as weights" (GY, 1988, p.3) and is a means of capturing the evolution of international prices of a basket of primaries.

The data for the six primary commodity price series (cocoa, coffee, copper, rice sugar, and tin) for the period 1900-2003 have been obtained from Pfaffenzeller (2007) and largely comprise data from the World Bank commodity price database up to and including 2003. Thereafter, the data comprises updated figures for the period 2004 to 2008. The annual average price for Thai 5% rice has been updated from the WB 2009 Pink Sheet, as have the updated figures for the International Coffee Organization's quote for Arabica coffee; the International Cocoa Organization's daily cocoa price; the International Sugar Agreement daily world price for sugar; and the London Metal Exchange prices for grade A copper, and tin (Pfaffenzeller *et al.*, 2007, p.3). Pfaffenzeller *et al.*'s (2007, pp.3-4) technical note provides a fuller description of all 24 non-fuel primary commodity prices which are incorporated into the GYCPI. Finally, each of the six prices series have been deflated by the annual rebased US consumer price index (1967=100, in accordance with DL). The graph of commodity prices used in this study is given in Figure 1 below:

[Figure 1 about here]

3.2 Updating world commodity production

Most of the production series are constructed using a combination of data provided in the appendix to DL (p. 306-309), for the period 1900-1987, and updates from various sources, namely the Food and Agricultural Organisation's (FAO) database and various reports and surveys from the International Cocoa Organisation (ICCO), the International Coffee Organisation (ICO), and the United States Geological Survey (USGS).

All of the food commodity production series comprise the data provided in the Appendix to DL (p. 306-9) and feature suitable alternative sources thereafter. The cocoa production data (measured in thousands of tonnes) originates from the DL data which has been provided by Gill and Duffus' serial *Cocoa Statistics* from 1901 to 1987. More recent data have been obtained from the ICCO's 2003/04 *Quarterly Bulletin of Cocoa Statistics* (QBCS) for the years until 1999[2]. The series has been constructed between 2000 up to and including 2004 using figures provided in the

[2] See http://www.icco.org/xls/Production_Table.jpg [Accessed 20 August 2009]

ICCO's 2004/05 *Annual Report*[3]. The data point for 2005 has been interpolated, while the 2006 and 2007 figures are revised estimates from online press releases associated with the latest ICCO QBCS[4]. DL provide coffee production data (measured in thousands of 60kg bags) from Brazil's Departamento Nacional do Cafes Annual Statistics, which spans 1930 to 1987. Thereafter, historical statistics have been provided by the ICO[5] until 2008. The DL dataset uses historical rice production data (measured in hundreds of millions of quintals) from FAO publications between 1904 and 1987 and this paper uses data from the FAO website[6] thereafter until 2007. Historical statistics on sugar crops production data (measured in hundreds of millions of quintals) have been made available to DL by FAO publications from 1903 to 1987. Post-1987 sugar production data has been obtained from the FAO website until 2007.

While DL use copper production statistics provided by the serial *Metallstatistik*, this paper uses a more complete and up-to-date series provided by the United States Geological Survey (USGS)[7], with measurements made in thousands of tonnes. The series on world copper mine production is a combination of all the available international data featured in both the *Mineral Resource in the United States* (MR) and the *Minerals Yearbook* publications. However, the USGS does note that the

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[3] See http://www.icco.org/about/anualreport.aspx [Accessed 20 August 2009]
[4] See http://www.icco.org/about/press2.aspx?Id=ox111384 [Accessed 20 August 2009]
[5] See http://www.ico.org/new_historical.asp At the time of writing, production data could be obtained by following the "Total production" link under "SUPPLY DATA" [Accessed 20 August 2009]
[6] See http://faostat.fao.org/site/567/default.aspx#ancor At the time of writing, rice production data could be accessed by selecting "World +", under "country"; "Rice, paddy" under "item"; and "Production quantity" under "element" for each year in the interactive database. For sugar production data, select "World +"; "sugar crops, nes"; and, "Production quantity" under the same headings in the interactive database [Accessed 20 August 2009].
[7] See http://minerals.usgs.gov/ds/2005/140/#data [Accessed 18 August 2009] At the time of writing, production data for copper and tin could be accessed by selecting the Microsoft Excel file under "supply-demand statistics" for copper and tin.

revisions to the data covering the period 1940-1985 have been made in order to incorporate new production techniques. The USGS is used as a source for world tin production data. While the DL data for the period 1900-1987 originates from various issues of the International Tin Council yearbooks, subsequent historical statistics have been obtained from the USGS website, where the production data represents the level of tin content within mine and mill production. The graph of commodity production used in this study is given in Figure 2 below:

[Figure 2 about here]

3.3 Updating annual world GDP

This paper does not use the annual world GDP series provided in the appendix of DL's (2003) work, which covers the period 1900-1987 (DL, 2003, p.306-9). However, this paper employs the same technique of constructing the annual world GDP series which originates from Maddison's work on historical statistics (1989, pp.111-3). The annual world GDP series is an aggregation of the total GDP (in millions of 1990 Geary-Khamis international dollars) for the OECD-16 countries, namely, Austria, Belgium, Denmark, Finland, France, Germany, Italy, Netherlands, Norway, Sweden, Switzerland, the United Kingdom, Australia, Canada, Japan, and the United States. Total GDP for each country is calculated by multiplying the annual population (in thousands) by annual GDP per capita (in 1990 Geary-Khamis international dollars), both of which can be retrieved from Maddison's online source[8].

[8] See http://www.ggdc.net/maddison/ [Accessed 10 August 2009]

In his technical note on the construction of historical statistics[9], Maddison emphasizes the greater importance of purchasing power parity (PPP) converters for inter-country comparisons of GDP levels over the use of exchange rates in excluding "the impact of inter-temporal price change". He cites the work of Kravis, Heston, and Summers (1982), who together initiated a sequence of phases of the International Comparison Programme (ICP) using a "highly sophisticated comparative pricing exercise in which national accounts expenditure in the participating countries in a given year was decomposed in great detail for representative items of consumption, investment and government services" (Maddison, 2001, p.171). Furthermore, the results of this exercise are then "multi-lateralized using the Geary-Khamis technique which ensured transitivity, base country invariance and additivity" (*ibid.*). This paper adheres to Maddison's use of 1990 as the benchmark year, rather than the more outdated 1980 international dollar measure used in DL. The graph of world income used in this study is given in Figure 3 below:

[Figure 3 about here]

4. Econometric Methodology

In order to ensure a full account of the temporal properties of the series in question, the unit root tests are conducted. These tests are presented in such a way as to assist the reader to draw equivalence with the theory presented in Section 3. Since DL

[9] See http://www.ggdc.net/maddison/Historical_Statistics/BackgroundHistoricalStatistics_09-s2008.pdf [Accessed 22 August 2009]

obtain a linear stationary process for commodity prices they apply the standard Dickey Fuller (1979) test as follows:

$$\Delta z_t = \alpha + \gamma z_{t-1} + \sum_{i=1}^{p} \Delta z_{t-i} + \varepsilon_t \qquad \varepsilon_t \sim \text{iid } N\left(0, \sigma^2\right) \tag{8}$$

Where z_t is the price series and p denotes the lag length selected according to the Schwartz Bayesian Criterion (SBC) to correct for plausible serially correlated errors. Since we propose a nonlinear model, this paper employs a test for unit roots under the null hypothesis against the alternative hypothesis of nonlinear Exponential Smooth Transition Auto Regressive (ESTAR) adjustment designed by Kapetanios, Shin and Snell (2003), (henceforth KSS2003). The test couches the theoretical underpinnings described in (4) by reformulating (8) to give the following nonlinear ESTAR process:

$$\Delta z_t = \gamma z_{t-1}\left[1 - \exp\left(-\theta z_{t-1}^2\right)\right] + \varepsilon_t \qquad \varepsilon_t \sim \text{i.i.d}\left(0, \sigma^2\right) \tag{9}$$

The null hypothesis of a unit root for this test procedure is $H_0 : \theta = 0$ against the alternative $H_1 : \theta > 0$. However, testing this null hypothesis directly is not feasible, since γ is not identified under the null. Thus, KSS2003 compute a Taylor series approximation to the ESTAR model under the null to obtain the following auxiliary regression:

$$\Delta z_t = \delta z_{t-1}^3 + \omega_t \tag{10}$$

In the general case, when the errors in (9) are serially correlated in a linear fashion then (9) may be extended to

$$\Delta z_t = \sum_{i=1}^{p} \beta_i \Delta z_{t-i} + \gamma z_{t-1}\left[1 - \exp\left(-\theta z_{t-1}^2\right)\right] + \varepsilon_t \qquad \varepsilon_t \sim \text{i.i.d}\left(0, \sigma^2\right) \tag{11}$$

and the auxiliary regression with p augmentations is obtained to be:

$$\Delta z_t = \delta z_{t-1}^3 + \sum_{i=1}^{p} \beta_i \Delta z_{t-i} + \omega_t . \tag{12}$$

where the lagged values of Δz_t are included to correct for plausible serially correlated errors. To choose the number of lags we follow the significance procedure proposed by KSS2003. In both cases, the null hypothesis to be tested is $H_0 : \delta = 0$ against the alternative $H_0 : \delta < 0$ using a $\hat{t}_{NL}$ test. KSS2003 show that the $\hat{t}_{NL}$ test does not have an asymptotic standard normal distribution and undertake stochastic simulations to obtain the asymptotic critical values.

Elliot *et. al.* (1996), ERS hereafter, derive a more powerful GLS detrending/demeaned based test in the context of linear unit root tests. In a more recent paper, Kapetanios and Shin (2008) extend the GLS detrending/demeaned procedure to test for unit roots allowing ESTAR adjustment concluding that the GLS based unit root tests are more powerful than the OLS alternative due to KSS2003. Following Kapetanios and Shin (2008) the GLS based unit root tests can be given by the following equation:

$$\Delta\tilde{z}_t = \delta\tilde{z}_{t-1}^3 + \sum_{i=1}^{p}\beta_i\Delta\tilde{z}_{t-i} + \omega_t \tag{13}$$

where $\tilde{z}_t = z_t - \hat{\mu}$ is the GLS based demeaned series and $\hat{\mu}$ is the GLS estimate. The associated t-statistic for the null hypothesis $H_0 : \delta = 0$ is denoted by t_{GLS-NL}. The associated critical values are obtained from Kapetanios and Shin (2008).

Based on the economic model derived earlier, [see equation (7)], one can motivate the specification and estimation of the following nonlinear Exponential Smooth Transition Auto Regressive Error Correction Model (ESTAR–ECM) due to Kapetanios, Shin and Snell (2006), [KSS (2006) hereafter].

$$\Delta x_t = \gamma\left[1 - \exp\left(-\theta z_{t-1}^2\right)\right] z_{t-1} + \beta\Delta y_t + u_t \tag{14}$$

where x_t denotes supply (production) and y_t denotes income; u_t is the error term. In order to overcome the problem that γ is not identified under the null, KSS (2006) propose that (14) is approximated using the Taylor series to obtain:

$$\Delta x_{1t} = \delta z_{t-1}^3 + \beta \Delta y_{2t} + u_t \tag{15}$$

From (15) we compute the $\hat{t}_{NLECM}$ statistic for the null of no cointegration, $H_0 : \delta = 0$, against the alternative of nonlinear ESTAR cointegration, that is, $H_0 : \delta < 0$. As with the $\hat{t}_{NL}$, KSS2006 undertake stochastic simulations to obtain the asymptotic critical values of $\hat{t}_{NLECM}$. KSS2003, KSS2006 and Kapetanios and Shin (2008) show, using Monte Carlo simulations, that the proposed tests, $\hat{t}_{NL}$, $\hat{t}_{NLECM}$ and t_{GLS-NL} respectively, have good size properties and superior power properties than the standard unit root or the Engle Granger test. The simulations conducted by KSS2006 show that the power gain is substantial, when θ is relatively small. They further demonstrate that the $\hat{t}_{NLECM}$ is superior to the standard Engle-Granger and $\hat{t}_{NLEG}$ when the regressors are weakly endogenous in the cointegrating regression.

In the analysis that follows, we apply both linear methods used by DL, and the nonlinear methods that we propose on the basis of the economic model suggested in Section 2.

5. Empirical Results and Policy Implications

Table 1 replicates the ADF procedure undergone in DL. It shows the results of the ADF tests carried out on the natural logarithm of the six commodity prices (after they have been deflated by the US consumer price index).[10]

[Table 1 about here]

DL show evidence that, using the ADF test with regards to commodity prices, the negative coefficient on the lagged dependent variable implies a tendency towards stationarity. When judged on those grounds, a similar result emerges from the ADF results of commodity prices in this paper. The fact that all commodity prices, except tin, within this paper are non-stationary $I(1)$ processes shows evidence which contradicts the crucial assumption of price stationarity within the Lewis model. DL criticize the low power of the ADF test and also mention certain caveats to the ADF results, namely "high autocorrelations at high frequency, coupled with a lack of long-run trend" (DL, p.299); secondly, the tests are unable to detect slow mean reversion which would otherwise be shown in variance-ratio statistics as calculated within a structural time series framework [outlined in Deaton and Laroque (1992)]. We acknowledge, here, the commonly-held idea that the ADF test has a tendency to under-reject the null hypothesis of a unit root. Consequently, this paper aims to address the issue by adopting the unit root tests due to KSS2003 in order to better

[10] The ADF tests are also carried out on the natural logarithm of commodity production for each of the six commodities, and the natural logarithm of world income. The tests are required for the subsequent cointegration analysis following (7). Though the results are not reported here for brevity, they are available from the authors on request.

characterize the price, production, and world income series used in this paper. The KSS2003 unit root test is applied based on the theoretical motivation illustrated in Section 4.

The results of the KSS2003 test are displayed in Table 1. We find that two commodities (copper and sugar) display a stationary process with ESTAR adjustment. This result is a marginal improvement over the DL model where only one commodity showed evidence of being a stationary process. Though the evidence is limited, one may conclude that the KSS2003 unit root tests lend support to the type of nonlinear adjustment that we describe in the theoretical model. Following KSS2003, although the t-ratio of θ does not provide a valid significance test, the 90% confidence intervals computed under the alternative hypothesis for θ do not include zero in two of the prices (copper and sugar) that are found to exhibit ESTAR adjustment. Finally, except for rice, all estimates of θ lie approximately within the interval [0.01,0.1] which is in line with the numerical experiments made by KSS2003 thereby implying an empirically meaningful range for θ.

Using the more advanced tests proposed by Kapetanios and Shin (2008), we conduct unit root tests using the GLS based approach. To facilitate comparison, we test for unit roots using the linear approach on demeaned variables due to the method by ERS, and then the ESTAR approach by Kapetanios and Shin (2008). The results are shown in Table 2.

[Table 2 about here]

This time the results lend stronger support to the theoretical underpinnings that we provide in Section 2. Using the linear approach based on the model by DL, we find evidence of two commodities (coffee and tin) being stationary. However, using the GLS-based demeaned approach due to Kapetanios and Shin (2008), we find that four (coffee, copper, sugar and tin) out of six commodities display stationary behavior with ESTAR adjustment. The estimates of θ and the associated standard error show that at the 95% confidence intervals computed under the alternative hypothesis for θ do not include zero in any of the prices except for rice. Finally, except for rice, all estimates of θ lie approximately within the interval [0.01,0.1] which is in line with the numerical experiments made by KSS (2003) thereby implying an empirically meaningful range for θ. The conclusion from these tests is that by using more powerful unit root tests we find strong evidence that nonlinear price adjustment exists as described by the theoretical underpinnings in Section 2.

Once more, the majority of price series are stationary with non-linear adjustment. Consequently, the results of this paper contrast significantly from those of DL, where the lower-power unit root tests showed that the majority of price series are non-stationary. Therefore the results of this paper provide more substantial evidence for DL's key assumption of price stationarity within the Lewis model framework.

Moving on to the cointegration tests between production and income, we employ the KSS2006 test for cointegration and compare the results with the linear cointegration test due to Engle and Granger (1987) for comparison. The results are shown in Table 3.

[Table 3 about here]

The results of the Engle Granger (1987) test for cointegration are listed in the second column of Table 3. Apart from copper and tin, none of the commodities display a cointegrating relationship between supply and income. The results do not change when testing for cointegration allowing for ESTAR adjustment using the KSS2006 approach. Table 3 shows the results of the KSS2006 approach of determining whether there is cointegration between commodity supply and world income in the third column. The test statistic is compared to the critical value at the 5% level in the case of cocoa, coffee, and sugar, which indicates the non-rejection of the null hypothesis of no cointegration for cocoa, coffee, rice and sugar. This implies that the same commodities, namely copper and tin, display cointegration with both linear and nonlinear adjustment. Finally the ESTAR-ECM model proposed by KSS2006 describes copper, tin and rice to display ESTAR adjustment. In general, using both linear and nonlinear (ESTAR) adjustment, the null of no cointegration between production and world income cannot be rejected for cocoa, coffee and sugar.

Following KSS2006, the $\hat{t}_{NLECM}$ test and the $\hat{t}_{NLEG}$ have good size and power properties compared to the $\hat{t}_{EG}$ test. The power gain from the $\hat{t}_{NLECM}$ test and the $\hat{t}_{NLEG}$ test over the $\hat{t}_{EG}$ test becomes substantial when $\hat{\theta}$ is relatively small. When the regressors are likely to be endogenous, the $\hat{t}_{NLECM}$ test tends to be more superior than the $\hat{t}_{NLEG}$. The results from Table 3 show that except for coffee, all the commodities generate an estimate which is quite small. The range of values lies within the interval [0.02, 0.09]. Overall, one can conclude that from the $\hat{t}_{NLECM}$ test results, which tend to be more superior in power and size properties, exactly half of the commodities considered show evidence of ESTAR adjustment. While two of these commodities (that is, copper and tin) are also found to display linear adjustment, we find that rice, which is found to show no meaningful long-run relationship (with linear adjustment) over time, exhibits ESTAR adjustment under the $\hat{t}_{NLECM}$ framework.

We can therefore deduce that while the cointegrating relationship between production and world income holds for copper, rice and tin, the adjustment to any disequilibrium is characterized as a nonlinear ESTAR process. In summary, we can say that the production of these three primary commodities will adjust at a faster rate to their long-run equilibrium as a result of changes in world income when the deviation of production from income is substantially large. For small deviations of the production and income relationship, the change in production would be limited or very small.

One may note that while four prices (that is, coffee, copper, sugar and tin) show nonlinear adjustment, this nonlinear adjustment in prices feeds into two commodities (that is, copper and tin) when analysing the cointegrating relationship between income and production. This may be a result of other factors affecting the adjustment process for the income-production relationship for rice.

The higher-power ESTAR-ECM due to KSS2006 show a more evidence of cointegration between the different production series and world income, compared to the Engle-Granger (1987) two-step procedure. With reference to the Lewis model, this paper argues that world production of one primary commodity is more likely to adjust to world income, and not vice versa, as world income levels in this study comprise the GDP of 16 countries, whereas the world production series focuses on only one primary commodity market at any one time.

The price mechanism within the Lewis model [see equation (2)] is implicit within the explanation of nonlinearity, as a rise in commodity prices (p_t) above the subsistence wage (p^*) under the Engle-Granger representation will have a positive and linear effect on commodity supply. In contrast, the adjustment of commodity supply under the nonlinear cointegration approach due to KSS2006 and ESTAR-ECM approach due to KSS2006 will depend on the extent of the commodity price rise relative to the subsistence wage. With no deviation of prices from the subsistence wage, supply in period t will be a function of supply in period (t-1) plus an error term. A large deviation in $\left(p_t - p^*\right)$ will incentivize supply to adjust to the long-run equilibrium, whereas a small deviation in prices will not, as a series of small adjustments would represent unsustainable costs to an economy. This provides the rationale for the nonlinear adjustment of commodity supplies to deviations in commodity prices. One may expect that this nonlinearity then feeds into the adjustment mechanism of the cointegrating relationship between income and production.

The nonlinear adjustment of log commodity production to log world income could be attributed to the fact that primary commodities take much longer to produce and as such, are more susceptible to government intervention when there are surges in world demand but scarcity of supply. For example, the surge in the world demand for rice in recent years prompted some rice-producing countries, such as India, China, Vietnam, and Egypt, to impose export restrictions on rice in order to preserve domestic rice supplies and to avoid domestic food price spirals in their respective countries (BBC, 2008)[11]. Such was the case concerning the world food crisis between 2007 and 2008, where "weather and crude oil price shocks resulted in contractions in food production in the wake of rising food demand brought about by rapid population growth in developing countries" (FAO, 2009, p.12). The recent crisis, where "demand factors (notably biofuel demand) were key" (*ibid.*), saw rice price volatility in the first four months of 2008 soar to levels which were five times higher than those in 2007 (FAO, 2009, p.11). At the time, the Thai government's decision to remain open to free trade in the face of export bans imposed by other major rice-producing countries was based

[11] See http://news.bbc.co.uk/1/hi/business/7328087.stm

on the country’s confidence in having sufficient domestic rice supplies[12], even when considering the large upwards deviation in world demand away from the long-run relationship. What is noteworthy is that this paper uses world income as a proxy, and that the use of Thai rice price series in the estimations makes this line of argument all the more appropriate. In contrast, the situation in other rice-producing countries, such as India (producing 10% of rice exports) and Vietnam (producing 15% of world rice exports)[13], demonstrates the nonlinear argument for cointegration. If the deviation of world demand away from long-run equilibrium had been small, then there would have been less incentive for these economies to enforce export bans as a way of preserving rice supplies in order to satisfy their domestic demand. This demonstrates the use economic events in explaining the nonlinearity observed.

6. Conclusion

This paper has sought to investigate how well a nonlinear model of commodity prices is supported by empirical evidence, using more advanced econometric techniques and more recent data on commodity prices and production, and world income. The model that we propose builds on the linear model proposed by DL by allowing for nonlinearities. This is carried out by modifying the supply equation of DL to allow for nonlinear adjustment and to this end employs a unit root test that allows for ESTAR adjustment. We observe that when applying the standard ADF test to all the commodity prices for our extended data set, we find similar results obtained by DL;

[12] See http://www.reuters.com/article/GCAAgflation/idUSSP24058020080501?pageNumber=1&virtualBrandChannel=0

[13] See http://www.unctad.org/infocomm/anglais/rice/ecopolicies.htm#pol

that is, barring one commodity, all of the primary commodity prices under observation are $I(1)$ processes. These results challenge one of the theoretical underpinnings of price stationarity within the DL framework of the Lewis model. However, when applying the Kapetanios and Shin (2008) GLS based unit root test that allows for nonlinear adjustment, we find that there is considerable evidence that commodity prices are stationary. Exactly two-thirds of the prices display a stationary process with ESTAR adjustment. Our results add support to the findings in a recent study by Balagtas and Holt (2009) who find significant evidence of nonlinear Smooth Transition Auto Regressive (STAR) dynamics in commodity prices. Further, this paper finds more evidence of cointegration between production and income when employing nonlinear tests for cointegration, due to KSS2006. We find support for cointegration with ESTAR adjustment for three commodities (that is, copper, rice and tin). The use of these more advanced econometric methods shows that the long-run relationship between world income and commodity production is better explained via nonlinearity rather than by the linear representation.

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Table 1: Linear and Nonlinear Unit Root Test Results (Prices)

	$\hat{t}_{ADF}$	$\hat{t}_{NL}$	$\hat{\theta}$	$t_{\hat{\theta}}$
Cocoa	– 2.59 (10)	– 2.19 (2)	0.052 [0.015]	3.47
Coffee	– 2.66 (0)	– 2.44 (0)	0.035 [0.011]	3.13
Copper	– 2.33 (11)	– 3.79 (1)**	0.048 [0.012]	3.86
Rice	– 1.15 (9)	– 1.26 (2)	0.015 [0.016]	1.92
Sugar	– 2.23 (2)	– 3.84 (6)**	0.083 [0.016]	4.89
Tin	– 2.98 (1)*	– 2.53 (1)	0.026 [0.009]	2.73

**, and * denote significance at the 1% and 5% levels respectively. The numbers in parentheses denote the lags chosen so that the error term is a white noise process. The numbers in square brackets denote the standard errors. The number of lags are chosen according to the General to Specific Methodology recommended by Kapetanios *et. al.* (2003).

Table 2: GLS Detrended Linear and Nonlinear Unit Root Test Results (Prices)

	$\hat{t}_{ERS}$	$\hat{t}_{GLS-NL}$	$\hat{\theta}$	$t_{\hat{\theta}}$
Cocoa	– 1.09 (2)	– 1.85 (0)	0.021 [0.0085]	2.49
Coffee	– 2.68 (0)**	– 2.23 (0)*	0.023 [0.0081]	2.88
Copper	– 1.93 (1)	– 2.97 (1)**	0.02 [0.009]	3.05
Rice	– 1.10 (2)	– 1.32 (1)	0.006 [0.004]	1.41
Sugar	– 1.82 (0)	– 2.90 (0)*	0.05 [0.014]	3.67
Tin	– 3.02 (1)**	– 2.54 (1)*	0.02 [0.009]	2.65

**, and * denote significance at the 1% and 5% levels respectively. The numbers in parentheses denote the lags chosen so that the error term is a white noise process. The numbers in square brackets denote the standard errors. The associated 95% critical value is –2.21.

Table 3: Cointegration Test and ESTAR – ECM Results (Production and Income)

	$\hat{t}_{EG}$	$\hat{t}_{NLEG}$	$\hat{t}_{NLECM}$	$\hat{\theta}$	$t_{\hat{\theta}}$
Cocoa	–2.99 (1)	–2.20 (1)	–2.16 (1)	0.02 [0.007]	3.01
Coffee	–2.24 (1)	–2.69 (1)	–2.79 (1)	0.25 [0.053]	4.67
Copper	–4.53 (1)**	–6.40 (1)**	–5.91 (1)**	0.06 [0.014]	4.76
Rice	–2.90 (1)	–3.18 (1)	–3.22 (1)*	0.05 [0.013]	4.33
Sugar	–2.30 (0)	–2.83 (0)	–2.81 (0)	0.05 [0.015]	3.51
Tin	–4.02 (0)**	–6.92 (1)**	–7.68 (1)**	0.09 [0.016]	5.75

***, **, and * denote significance at the 1%, 5%, and 10% levels respectively. The numbers in parentheses denote the lags chosen so that the error term is a white noise process. The numbers in square brackets denote the standard errors.

Figure 1

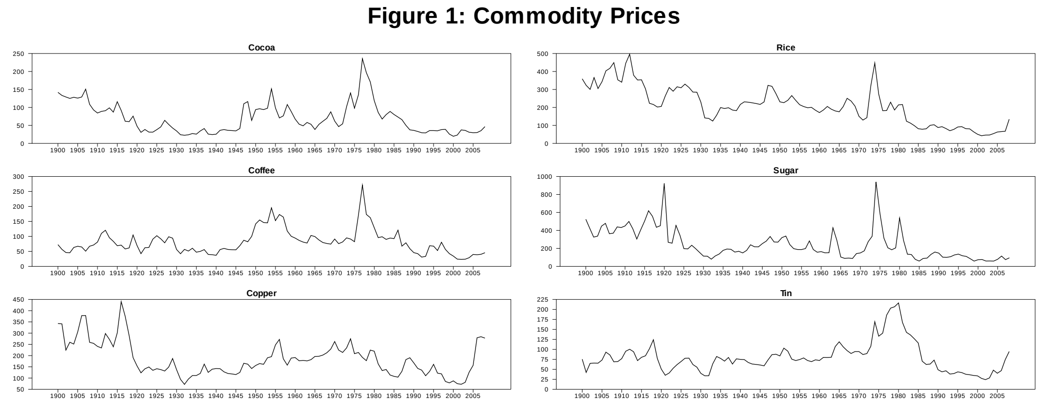

Figure 2

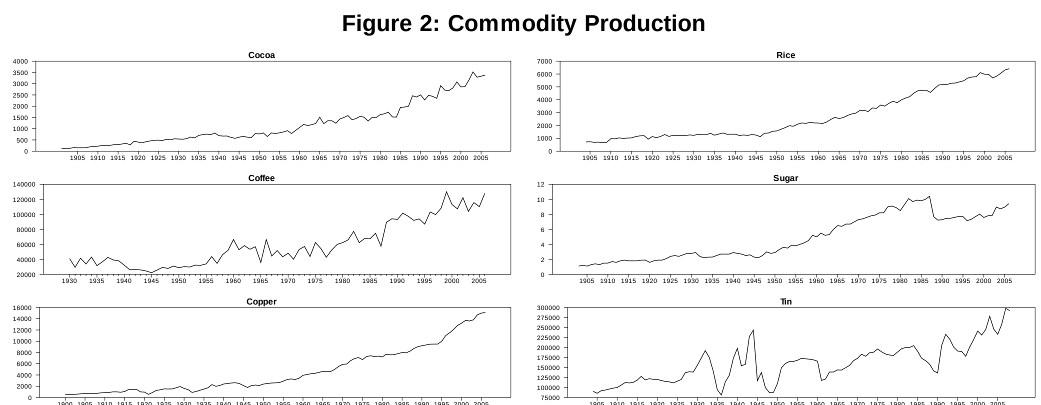

Figure 3

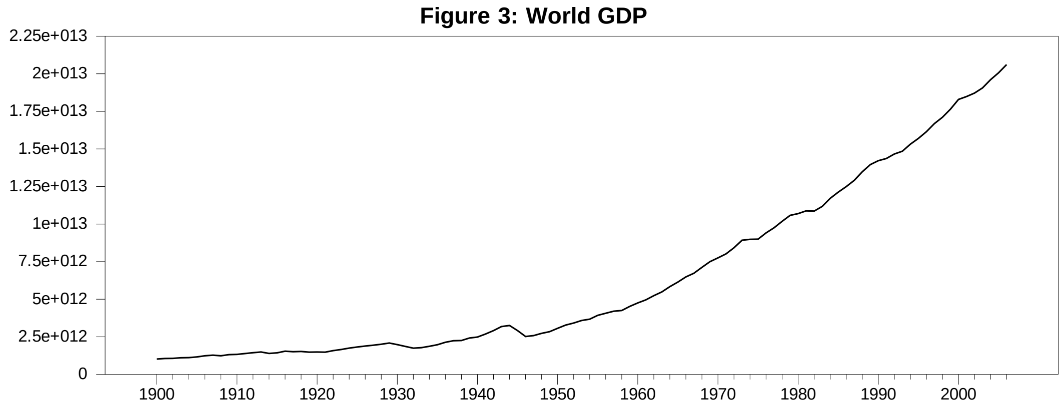

Figure 3: World GDP
2.25e+013
2e+013
1.75e+013
1.5e+013
1.25e+013
1e+013
7.5e+012
5e+012
2.5e+012
0
1900
1910
1920
1930
1940
1950
1960
1970
1980
1990
2000